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US MONETARY POLICY 1988–2004: AN EMPIRICAL ANALYSIS

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ABSTRACT: Relationships between the Federal funds rate, unemployment, inflation, and the long-term government-bond rate are investigated with cointegration techniques. We find a stable long-term relationship between the Federal funds rate, unemployment, and the bond rate. This relationship is interpretable as a policy target because deviations are corrected primarily via the Federal funds rate. A traditional Taylor-type rule is clearly rejected by the data. Inflation does thus only influence the instrument indirectly via the bond rate, but we find that inflation is controllable with the Federal funds rate. The results are in accordance with recent developments in monetary theory stressing management of expectations as an important transmission channel.

KEYWORDS: Cointegration; Equilibrium Correction; Monetary Policy; Taylor rule; Bond rate.

JEL CLASSIFICATION: C32; E52.

1 INTRODUCTION

In his seminal contribution, Taylor (1993) suggests that the complex monetary policy process in the US 1987 – 1992 can be summarized by a simple policy rule, in which the Federal funds rate reflects deviations of inflation and activity from their policy targets. Over the past decade, several authors have elaborated on the so-called Taylor-rules by introducing interest-rate smoothing; discussing the the role of expectations, forecasts, and data revisions; and analyzing the robustness of the policy rules to different measures of

The authors would like to thank Christopher Bowdler, Eilev Jansen, Henrik Jensen, Søren Johansen, Katarina Juselius, Dan Knudsen, and Hans Christian Kongsted for comments. Remaining errors and shortcomings are the sole responsibilities of the authors. The empirical analysis was carried out using Ox, see Doornik (2002), and PcGive, see Doornik and Hendry (2001).

activity and inflation, see *inter alia* Judd and Rudebusch (1998), Evans (1998), Clarida, Galí, and Gertler (1998), Clarida, Galí, and Gertler (2000), Carey (2001), Orphanides (2001), Ball and Tchaidez (2002), and Orphanides and Williams (2003).

In this paper we scrutinize the empirical regularities between activity (as measured by the monthly unemployment rate), inflation, a 10-year government bond rate, and the Federal funds rate for the period since Greenspan became chairman for the Federal reserve board: 1988 : 1 – 2004 : 8. Our main goal is to analyze the existence of a monetary policy rule in terms of observable variables, and using a cointegrated vector autoregressive (VAR) system we distinguish between longer-term targets and shorter-term dynamic adjustment. In this line of thoughts, the actual reaction function represents the dynamic adjustment towards equilibrium making interest rate smoothing an integral part of the model, and we argue that a long-term relationship can only be interpreted as a policy rule if deviations from the implied target explain changes in the Federal funds rate. In this way we do not merely estimate the coefficients of a postulated policy rule, we also test implications of the identification scheme. In the applied system approach the dynamic adjustment of all variables can be characterized simultaneously, and if a stable monetary policy rule is found, we can analyze the effectiveness of the instrument in the control of the policy objectives.

As a second contribution we use the cointegrated VAR model directly to approximate the evolution of the natural rate of unemployment. The natural rate is approximated by a high degree polynomial in time and the shape of the polynomial is determined simultaneously with the VAR coefficients, exploiting all the information in the sample. This is contrary to most other research, where the natural rate is determined outside the model using a polynomial or a linear filter, while the econometric models are estimated conditional on the natural rate.

A third contribution of the paper is the introduction of the bond rate, which allows us to characterize the interaction between monetary policy and market interest rates. Several authors have advocated for the important role of the bond rate for monetary policy, see *inter alia* Bernanke and Blinder (1992), Goodfriend (1991), and Goodfriend (1998). In addition, there has been an increasing interest in the information content of the yield-curve for future inflation, future activity, or future short and long rates, cf. Mishkin (1990), Estrella and Hardouvelis (1991), Hardouvelis (1994), and the survey by Estrella and Mishkin (1996). According to visual inspection the bond rate seems to be of importance in empirical research of US monetary policy, cf. Christensen (2002), and from a theoretical perspective there has been a growing interest in models including the long rates, cf. Ang, Dong, and Piazzesi (2004), Hördahl, Tristani, and Vestin (2004), and Piazzesi (2005) to mention a few important contributions. According to our empirical results the inclusion of the bond rate is crucial. In particular, we find a stable long-term relationship between the Federal funds rate, unemployment, and the bond rate; and a significantly equilibrium correction of the Federal funds rate allows for a policy-rule interpretation. By excluding the bond rate we find a new long-term relationship with the

inflation rate included, but the results are unstable and the Federal funds rate does not adjust to narrow the gap between the actual and the equilibrium rate; a traditional Taylor rule is thus clearly rejected by the data. We interpret this rejection as an indication that the bond rate contains relevant information on the inflation process besides information on financial market conditions in general.

We are of course fully aware that the monetary policy process in real-time is far more complicated than a simple policy rule suggests. Nevertheless, the present study gives a parsimonious but still empirically relevant representation of the kind of factors that have entered the decision making process. Woodford (2003) makes a distinction between central banking as setting a short-term interest rate and central banking as management of expectations, and Kohn and Sack (2003) find that statements by the Federal Open Market Committee, FOMC, and testimonies in Congress by Chairman Greenspan significantly affect market interest rates at longer maturities. A straightforward way to think of this result is that the monetary policy objectives as well as FOMC's assessment of the actual cyclical developments are clarified via testimonies and statements. Accordingly market interest rates can move in advance of the Federal funds rate in case the perceptions of market participants change, and our findings regarding the role of the bond rate can easily be interpreted as the outcome of such a process with forward-looking expectations.

The rest of the paper is organized as follows. First, Section 2 presents a simple theoretical framework for the analysis. Section 3 and 4 then presents, respectively, the data measurements and the econometric approach, while Section 5 reports the empirical results. Finally, Section 6 concludes.

2 A SIMPLE FRAMEWORK FOR TAYLOR RULES

In an important paper, Taylor (1993) suggested that the FOMC has managed the Federal funds rate according to the simple linear formula

$$f_t = \pi_t + \lambda_1 \cdot (u_t - u_t^*) + \lambda_2 \cdot (\pi_t - \pi_t^*) + \kappa_0,$$

where f_t denotes the Federal funds rate, π_t and π_t^* denote the inflation and the monetary policy target, u_t and u_t^* denote the unemployment rate and the natural rate of unemployment, and the constant term, κ_0 , is interpretable as the target real interest rate in equilibrium. If the inflation target is assumed constant, as it is usually the case, π^* is not empirically identifiable and the relation collapses to

$$f_t = \lambda_1 \cdot (u_t - u_t^*) + (1 + \lambda_2) \cdot \pi_t + \kappa_1, \quad (1)$$

where $\kappa_1 = \kappa_0 - \lambda_2 \pi^*$.

In the basic formulation, the relation (1) is contemporaneous, and some empirical applications use (1) as a basis for a linear regression, see e.g. Evans (1998) and Ball and Tchaidze (2002). Alternatively a partial adjustment structure can be considered, e.g.

$$f_t = \rho \cdot f_{t-1} + (1 - \rho) \cdot [\lambda_1 \cdot (u_t - u_t^*) + (1 + \lambda_2) \cdot \pi_t + \kappa_1], \quad (2)$$

where $\rho > 0$ implies interest rate smoothing, see Judd and Rudebusch (1998), Orphanides (2001), and English, Nelson, and Sack (2003).

Empirically, time series for inflation and interest rates are highly persistent, and they are often well approximated as unit root processes. This complicates inference on the parameters in (1) due to the spurious regression problem, and a natural starting point for an empirical analysis is therefore formal tests for integration and cointegration. If the variables in (1) are difference stationary, or I(1), a convenient representation of the dynamics is the (conditional) equilibrium correction form

$$\Delta f_t = \gamma_1 \cdot \Delta(u_t - u_t^*) + \gamma_2 \cdot \Delta\pi_t + \alpha \cdot [f_{t-1} - \lambda_1 \cdot (u_{t-1} - u_{t-1}^*) - (1 + \lambda_2) \cdot \pi_{t-1} - \kappa_1].$$

This is a natural generalization of the model in (2) and it formally redefines the Taylor rule from a contemporaneous reaction function to a longer-term target value for the policy rate, $f_t^* = \lambda_1 \cdot (u_{t-1} - u_{t-1}^*) + (1 + \lambda_2) \cdot \pi_{t-1} + \kappa_1$. For f_t^* to be interpretable as a target value for monetary policy it must hold that the Federal funds rate, f_t , corrects deviations from the target, $f_t - f_t^*$, which implies that α should be significantly negative. Within the vector autoregressive framework that we adopt in the empirical analysis in Section 5 we can formally test hypotheses on the equilibrium correction structure, and therefore also test whether f_t^* can be interpreted as a policy target.

2.1 THE ROLE OF EXPECTATIONS

Several authors have stressed the forward looking nature of monetary policy and emphasized the role of expectations, see Svensson (2003) for a review. Due to the considerable time lag in monetary policy it makes sense for the monetary authority to be preemptive and act on expected future conditions. Forward looking Taylor rules have been estimated in *inter alia* Clarida, Galí, and Gertler (1998), Clarida, Galí, and Gertler (2000), Orphanides (2001), and Orphanides and Williams (2003). They usually have the form

$$f_t = \lambda_1^* \cdot E[(u_{t+h} - u_{t+h}^*) | \mathcal{I}_t] + (1 + \lambda_2^*) \cdot E[\pi_{t+h} | \mathcal{I}_t] + \kappa_1^*, \quad (3)$$

where $E[\cdot | \mathcal{I}_t]$ denotes the expectation conditional on the information set at time t , and h denotes the forecast horizon. Estimation is typically based on rational-expectations-type moment conditions involving instruments in the information set, \mathcal{I}_t .

In a linear (or linearized) model of the economy, the rational expectations are themselves linear functions of realized values of the variables in the information set, i.e.

$$\begin{aligned} E[u_{t+h} - u_{t+h}^* | \mathcal{I}_t] &= \delta_{u,u} \cdot (u_t - u_t^*) + \delta_{u,\pi} \cdot \pi_t \\ E[\pi_{t+h} | \mathcal{I}_t] &= \delta_{\pi,u} \cdot (u_t - u_t^*) + \delta_{\pi,\pi} \cdot \pi_t, \end{aligned}$$

possibly including more dynamics, see Orphanides and Williams (2003) and Ireland (2003). Substituting the forecast functions into (3) yields a representation which is observational equivalent to (1), with $\lambda_1 = \lambda_1^* \cdot \delta_{u,u} + (1 + \lambda_2^*) \cdot \delta_{\pi,u}$ and $(1 + \lambda_2) = \lambda_1^* \cdot \delta_{u,\pi} + (1 + \lambda_2^*) \cdot$

$\delta_{\pi,\pi}$. Orphanides and Williams (2003) denote the resulting Taylor rule *outcome-based*, and it is interpretable as a reduced form representation of the entire policy process in terms of observables. This is sufficient for measuring the stance of monetary policy, describing developments in the monetary policy process, and predicting future interest rate movements. In practice, of course, expectations need not to be rational, and the reduced form is compatible with other configurations of expectation formations.

Besides this general treatment of expectations, we want specifically to analyze the role of financial market information. As noted by Goodfriend (1991) and Goodfriend (1998) the long-term bond yield automatically moves with inflation expectations of the agents in financial markets. This information could be useful for the monetary authority. Either directly as an explanatory variable in a structural Taylor-rule, or as an instrument in the forecasting equation, implying a bond-yield in the outcome-based rule for the entire policy process. The bond rate could also contain other relevant information; a sudden increase in the bond rate could reflect a declining credibility of monetary policy and the FOMC could react by a preemptive increase in the Federal funds rate. In theory the relationship between the short- and the long rate is normally turned the other way round, making long-term rates a function of expected future short-term rates as in the expectation theory of the term structure, even if Schiller (1990) acknowledges that a lot of evidence speaks against the empirical validity of the theory. Note, however, that a causation in both directions is not a problem for the current empirical analysis, as the statistical model allows all variables to be endogenous from the outset.

Bernanke and Blinder (1992) also use the spread between the Federal fund rate, f_t , and a long term bond rate, b_t , as an indicator of the stance of monetary policy and estimates reaction function for $f_t - b_t$ as well as for f_t . Mehra (2001) and Carey (2001) include the bond rate as an additional variable in a Taylor rule. The use of the term structure of interest rates also fits with recent research on the spread between the Federal funds rate and a bond yield as a predictor of future inflation or activity, cf. Mishkin (1990), Estrella and Hardouvelis (1991), and Estrella and Mishkin (1996).

Since b_t will react with a one-to-one impact from inflation expectations, and inflation is already present in the target (1), we insert only the 'new' information as measured by the real bond rate, $b_t - \pi_t$, and correct for the average 'tilt' of the yield curve, τ , to obtain

$$f_t = \lambda_1 \cdot (u_t - u_t^*) + (1 + \lambda_2 - \lambda_3) \cdot \pi_t + \lambda_3 \cdot b_t + \kappa_2,$$

where $\kappa_2 = \kappa_0 - \lambda_2\pi^* - \lambda_3\tau$. If there is a one-to-one impact from the bond rate to the Federal funds rate, $\lambda_3 = 1$, we obtain a simple Taylor-type target for the interest rate spread:

$$f_t = b_t + \lambda_1 \cdot (u_t - u_t^*) + \lambda_2 \cdot \pi_t + \kappa_2. \quad (4)$$

In the empirical analysis, we consider relations of the form (1) and (4) as candidates for the longer-term policy target in terms of observables.

3 DATA MEASUREMENTS

To analyze the empirical evidence on interest rate setting and the role of monetary policy, we consider a data vector, $Y_t = (f_t : b_t : u_t : \pi_t)'$, comprising the effective Federal funds rate, f_t ; a 10 year constant maturity Treasury bond rate, b_t ; the unemployment rate, u_t , calculated as the number of unemployed as a percentage of the civilian labour force (both seasonally adjusted); and core inflation measured as 100 times the year-on-year change in the log transformed consumer price index excluding food and energy, π_t . The considered effective estimation sample is 1988 : 1 – 2002 : 12, and we condition the analysis on the last months of 1987. Observations for 2003 : 1 – 2004 : 8 are used for out-of-sample analysis.

The time series are presented in Figure 1. Graph (A) depicts the Federal funds rate and the Treasury bond rate while graph (B) depicts the spread, $f_t - b_t$. The interest rates have some similarities, but the fluctuations in the Federal funds rate, f_t , are larger than those in the bond rate, b_t , and on average f_t has been lower than b_t .

Graph (C) depicts the unemployment rate. The sample period covers a slack in the early 1990s and a subsequent long upturn ending sharply in 2000, where the unemployment rate is down to 4%. A comparison of (C) and (A) indicates a negative correlation between u_t and f_t , and there is also a clear correlation between u_t and the interest rate spread, $f_t - b_t$, in (B). Some authors have suggested that a considerable proportion of the fall in the unemployment rate over the period can be attributed to a fall in the natural rate of unemployment, u_t^* , see *inter alia* Ball and Tchaidez (2002). The natural rate is unobservable, but the considered system of variables may contain information on the changes in u_t^* . To extract this information we assume that changes in the natural rate are smooth and we use a fifth-order polynomial in the time index t to approximate the natural rate, i.e. $u_t^* = g(t)$, where $g(t)$ is a polynomial. In doing so, we reformulate the system in terms of the unemployment gap, $u_t - u_t^*$.

Finally graph (D) depicts core inflation. Inflation has been steadily decreasing over the period with bouts of rising inflation in early 1990s and early 2000s. The downward trend in π_t is also reflected in decreasing nominal interest rates, whereas high frequency movements seem less related.

4 ECONOMETRIC APPROACH

As a statistical framework for the analysis we use the cointegrated vector autoregressive (VAR) model of Johansen (1988), Johansen (1991) and Johansen (1996). It is important to note that by adopting this framework we do not assume unit roots in the variables. On the contrary, the starting point is a *stationary* VAR model and unit roots impose testable restrictions on the dynamic system.

The starting point is the p -dimensional VAR model in equilibrium correction form:

$$H(r) : \Delta Z_t = \alpha (\beta' Z_{t-1} + \mu') + \sum_{i=1}^{k-1} \Gamma_i \Delta Z_{t-i} + \epsilon_t, \quad t = 1, 2, \dots, T. \quad (5)$$

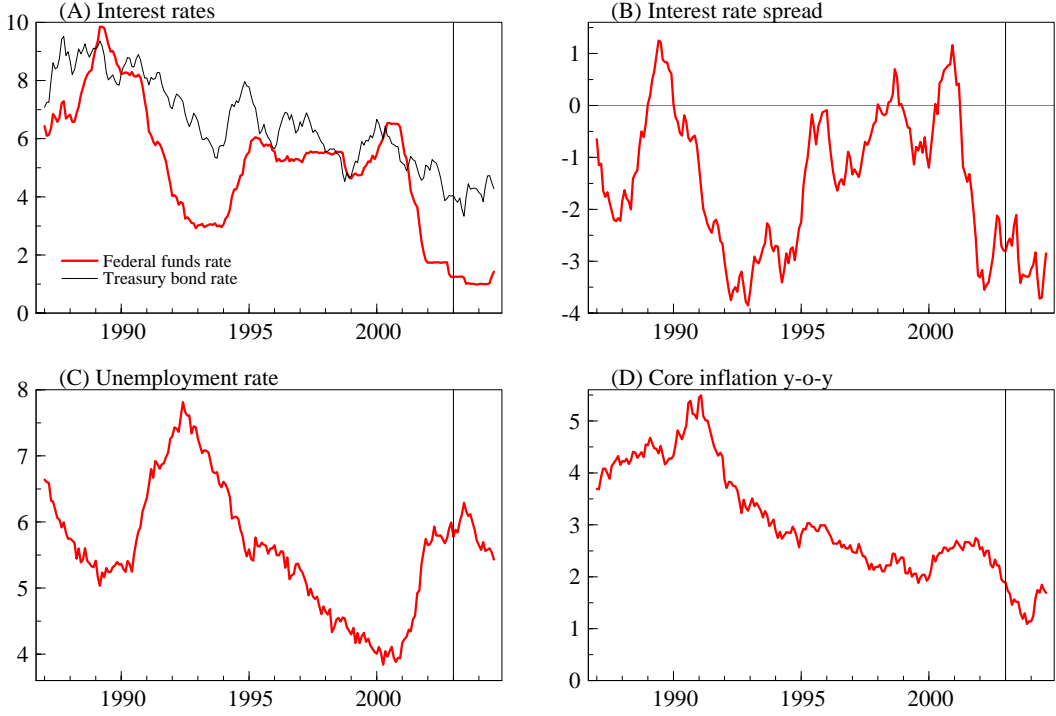


Figure 1: The data 1987 : 1 – 2004 : 8. The vertical lines indicate the end of the estimation sample 2002 : 12.

Here α and β are $p \times r$ matrices where $\beta' Z_t$ represents the $r \leq p$ cointegrating relationships and α gives the direction and speed of adjustment towards equilibrium. The remaining autoregressive coefficients, $\Gamma_1, \dots, \Gamma_{k-1}$, are of dimension $p \times p$, and μ is an $r \times 1$ constant term. Throughout, we assume that ϵ_t is an i.i.d. Gaussian sequence, $N(0, \Omega)$, and we condition on the initial values, Z_{-k+1}, \dots, Z_0 , see Johansen (1996, chapter 6).

Under the assumption that the characteristic polynomial, $A(z) = (1 - z)I - \alpha\beta'z - \sum_{i=1}^{k-1} \Gamma_i (1 - z)z^i$, has $p - r$ unit roots and the remaining roots are all located outside the unit circle, the variables, Z_t , have the representation

$$Z_t = C \sum_{i=1}^t \epsilon_i + C(L) \epsilon_t + \tau_0, \quad (6)$$

where $C = \beta_{\perp} (\alpha'_{\perp} (I - \Gamma_1 - \dots - \Gamma_{k-1}) \beta_{\perp})^{-1} \alpha'_{\perp}$ is a $p \times p$ dimensional long-run impact matrix of rank $p - r$, $C(L)$ is a convergent polynomial, and τ_0 depends on μ and the initial values, see Johansen (1996, Theorem 4.2). It follows from (6) that the non-stationarity of Z_t is driven by $p - r$ common stochastic trends defined as $\alpha'_{\perp} \sum_{i=1}^t \epsilon_i$, with loadings in the individual variables given by the first part of the matrix C . The interpretation of a coefficient C_{ij} in C is the long-run effect on variable i from an innovation to ϵ_j .

Johansen and Juselius (2001) use this insight to discuss the issue of *controllability* of a target variables with a given instrument in the cointegrated VAR model. A target,

$d'Z_t$, is said to be controllable with an instrument $a'Z_t$, where a and d are p -dimensional vectors, if $d'Y_t$ can be made stationary around a target value d^* by intervening in $a'Z_t$ at all points in time. The necessary condition for controllability is therefore that a shock to the instrument has a non-zero long-run impact on the target, i.e. that $d'Ca \neq 0$. As an example, let $Z_t = (f_t : b_t : u_t : \pi_t)'$ as in our empirical analysis in Section 5. A necessary condition for controllability of the inflation rate, π_t , with the Federal funds rate, f_t , is that $C_{41} \neq 0$. This is a testable hypothesis in the cointegrated VAR model.

4.1 DETERMINISTIC SPECIFICATION

In the specification in (5) the constant term is restricted to be proportional to α , which produces non-zero means in the long-run relationships but no deterministic linear trends in the variables. This is our preferred specification as it may be hard from economic theory to rationalize an autonomous drift in interest rates and inflation. Most of the time series in Figure 1 show a trending behavior in the sample period, however, and from an empirical point of view deterministic linear trends may be a reasonable alternative to stochastic trends. In the present paper we therefore take a pragmatic approach and conduct a sensitivity analysis where we also present results allowing for linear trends in all variables.

To estimate the natural rate of unemployment, $u_t^* = g(t)$, where $g(t)$ is a polynomial in the time index t , we formulate a cointegrated VAR model with additive corrections, see Nielsen (2004b). In particular, we let Z_t in (5) be a vector of unobserved variables, which is related to the observations Y_t by the additive equation

$$Z_t = Y_t - \theta D_t, \quad (7)$$

where D_t is a $n \times 1$ vector of deterministic variables and θ is a matrix of coefficients. In the present case $D_t = (t : t^2 : \dots : t^5)'$ so that θD_t is a fifth-order polynomial. To avoid polynomial terms in f_t , b_t , and π_t , we impose restrictions on θ of the form

$$Z_t = \begin{pmatrix} f_t \\ b_t \\ u_t - u_t^* \\ \pi_t \end{pmatrix} = Y_t - \theta D_t = \begin{pmatrix} f_t \\ b_t \\ u_t \\ \pi_t \end{pmatrix} - \begin{pmatrix} \theta_{11} & 0 & 0 & 0 & 0 \\ \theta_{21} & 0 & 0 & 0 & 0 \\ \theta_{31} & \theta_{32} & \theta_{33} & \theta_{34} & \theta_{35} \\ \theta_{41} & 0 & 0 & 0 & 0 \end{pmatrix} \begin{pmatrix} t \\ t^2 \\ t^3 \\ t^4 \\ t^5 \end{pmatrix}. \quad (8)$$

Our preferred specification excludes linear trends in f_t , b_t , and π_t by setting $\theta_{11} = \theta_{21} = \theta_{41} = 0$. The alternative specification with deterministic linear trends in all variables leaves θ_{11} , θ_{21} , and θ_{41} unrestricted.

To estimate the cointegrated VAR system in (5) with the additive corrections in (7), we apply the switching algorithm proposed in Nielsen (2004a). First, conditional on an estimate $\tilde{\theta}$ of θ , the parameters in (5) can be estimated in a standard cointegration analysis for the corrected variables, $Z_t = Y_t - \tilde{\theta} D_t$. In the next iteration we can find

an updated estimate $\hat{\theta}$ of θ conditional on the estimated VAR parameters. In particular, the estimated residuals are given by, $\hat{e}_t = \hat{A}(L) Y_t - \hat{\alpha} \hat{\mu}'$, where $\hat{A}(L)$ is the estimated characteristic polynomial. Under the model specified in (5) and (7), the residuals can be written as

$$\hat{e}_t = \hat{A}(L) \theta D_t + \epsilon_t = \hat{H}_t \text{vec}(\theta) + \epsilon_t,$$

where $\text{vec}(\theta)$ stacks the columns of θ , and $\hat{H}_t = D_t' \otimes \hat{A}(L) = (\hat{A}(L) D_{1t} : \hat{A}(L) D_{2t} : \dots : \hat{A}(L) D_{nt})$. In this model θ can be found as the GLS type estimator

$$\text{vec}(\hat{\theta}) = \left(\sum_{i=1}^T (\hat{H}_i' \hat{\Omega}^{-1} \hat{H}_i) \right)^{-1} \left(\sum_{i=1}^T (\hat{H}_i' \hat{\Omega}^{-1} \hat{e}_i) \right), \quad (9)$$

see also Tsay, Peña, and Pankratz (2000). Full-information maximum likelihood estimates can be obtained by iterating between the two steps until convergence, and the restrictions in (8) can easily be imposed in the GLS step (9).

5 EMPIRICAL ANALYSIS OF US MONETARY POLICY

In this section we look at the empirical evidence on the interdependencies between interest rates, inflation and unemployment and their implications for monetary policy.

As a starting point we consider the unrestricted model $H(4)$ under the additive specification in (8) as a representation of the autocovariance structure in the data. Information criteria and successive testing for removal of lags indicate that a lag length of $k = 3$ is sufficient to capture the dependence in the time series, and we choose this model in the following.

Table 1 reports a battery of misspecification tests. Overall the model is well behaved. The only deviation from the null hypotheses of a well specified model is the rejection of Gaussian residuals in the equation for inflation, due to two outliers in the year 1990. One purpose of this paper is to analyze the policy-response to inflation shocks, and therefore we prefer not to remove the shocks by including dummy variables. We can add that other choices of lag-length as well as the presence of the two dummies only marginally change the results presented below.

5.1 RANK DETERMINATION

Based on the statistical model we turn to the number of unit roots. The cointegration models form a nested sequence, $H(0) \subset \dots \subset H(r) \subset \dots \subset H(p)$. The number of unit roots, $p - r$, and the number of long-run relations, r , can be determined by likelihood ratio (LR) tests in the sequence. Here we consider the hypotheses $H(r) \mid H(p)$, parallel to the well-known trace tests.

The results of the LR tests are reported in Table 2 for the two deterministic specifications. The model $H(0)$ with four unit roots and no cointegration is clearly rejected. The model $H(1)$ with $r = 1$ long-run relationship and three unit roots has p -values of 11%

	AR(1)	AR(1-7)	ARCH(1-4)	Normality
Δf_t	3.076 [0.08]	6.335 [0.50]	5.823 [0.21]	3.340 [0.19]
Δr_t	2.422 [0.12]	12.023 [0.10]	4.525 [0.34]	4.726 [0.09]
Δu_t	0.508 [0.48]	5.986 [0.54]	0.753 [0.94]	0.553 [0.76]
$\Delta \pi_t$	0.052 [0.82]	5.227 [0.63]	1.498 [0.83]	10.914 [0.00]
Multivariate tests	2.104 [0.99]	119.133 [0.30]	414.690 [0.30]	20.811 [0.01]

Table 1: Tests for misspecification of the unrestricted VAR(3). Figures in square brackets are p -values. AR(1) are LM tests for first order autocorrelation in the residuals and are distributed as $\chi^2(1)$ and $\chi^2(16)$ for the single equation and multivariate tests respectively. AR(1-7) are LM tests for autocorrelation up to seventh order and are distributed as $\chi^2(7)$ and $\chi^2(112)$ respectively. ARCH (1-4) are LM tests for ARCH effects up to fourth order and are distributed as $\chi^2(4)$ and $\chi^2(400)$ respectively. The last column reports the Doornik and Hansen (1994) tests for normality, distributed as $\chi^2(2)$ and $\chi^2(8)$ respectively.

	(A) Fifth order polynomial in u_t				
	$H(0)$	$H(1)$	$H(2)$	$H(3)$	$H(4)$
Log-likelihood	1389.82	1407.38	1420.14	1425.58	1427.74
LR: $H(r) H(4)$	75.83 [0.00]	40.71 [0.11]	15.21 [0.21]	4.32 [0.37]	...

	(B) Fifth order polynomial in u_t and linear trends in $(f_t : b_t : \pi_t)$				
	$H(0)$	$H(1)$	$H(2)$	$H(3)$	$H(4)$
Log-likelihood	1391.42	1408.88	1421.49	1432.19	1438.06
LR: $H(r) H(4)$	93.27 [0.00]	58.35 [0.07]	33.14 [0.17]	11.73 [0.41]	...

Table 2: Likelihood ratio tests for rank determination, $H(r)|H(p)$. Asymptotic p -values in square brackets. The asymptotic distributions have been simulated for the current deterministic specifications.

and 7% for the two different deterministic specifications, indicating the presence of one stationary relation. In the models with $r = 1$, the largest unrestricted eigenvalues in the companion matrix have moduli given by 0.89 in both the two deterministic specifications, indicating that the adjustment towards equilibrium is relatively slow. In the larger model $H(2)$ the largest unrestricted eigenvalues have moduli of 0.95 and 0.92 in the two cases, which again seem to reflect the non-stationarity of the second relation. In the main set of results presented below we therefore take the model $H(1)$ as the preferred, but in Section 5.6 we discuss the possible interpretation of a second long-run relationship in the system.

In the preferred model $H(1)$ the linear trends are largely insignificant in the interest rates and inflation and a LR test for the hypothesis $\theta_{11} = \theta_{21} = \theta_{41} = 0$ give a statistic of $-2 \cdot (1407.38 - 1408.88) = 3.0$, which is not significant in a $\chi^2(3)$.

5.2 LONG-RUN STRUCTURE

Taking the model $H(1)$ with $r = 1$ long-run relation as the preferred, the unrestricted estimates of α and β^* are reported in Table 3 under \mathcal{H}_0 . Panel (A) again reports the results for the preferred specification, while panel (B) reports the results for the alternative specification with linear trends in f_t , b_t , and π_t . The long-run relation is normalized on the Federal funds rate, and the coefficient to the Treasury bond rate is around minus unity and statistically significant. The coefficient to the unemployment gap, $u_t - u_t^*$, is just below three, indicating that labour market pressure is associated with a high Federal funds rate. The coefficient to core inflation is positive as well, the opposite of the expected for a monetary policy rule.

For the long-run relation to be interpretable as a monetary policy rule, a necessary condition is that the adjustment coefficient in the Federal funds equation, α_1 , is significantly negative, such that deviations from equilibrium are corrected by policy actions. According to the t -ratios for α this is indeed the case with t -ratios of -3.65 and -4.36 for the two deterministic specifications, respectively. There is also a significantly negative impact in the inflation equation, indicating that a high interest rate relative to the target lowers inflation. The adjustment in the equation for the bond rate is close to zero and the adjustment in the unemployment equation is also insignificant.

Based on the unrestricted coefficients, the long-run relation looks like a monetary policy target. A conventional Taylor rule would imply a zero coefficient for the Treasury bond rate, $\beta_2 = 0$. Imposing this restriction gives the results reported under \mathcal{H}_1 . Under this hypothesis the presence of linear trends in the variables is important, and the results differ markedly between the two specifications. In panel (A), without linear trends in the variables, the coefficient to the inflation rate is significantly positive, giving no support for a traditional Taylor-rule as a long-run target for monetary policy. This is also emphasized by the fact that under the restriction in \mathcal{H}_1 the Federal funds rate do not adjust to deviations from the target. In panel (B), allowing for deterministic linear trends in the variables, the coefficient to inflation is significantly negative, indicating a possible Taylor rule. Interestingly, however, the Federal funds rate, f_t , adjusts away from equilibrium, invalidating the interpretation of the relation as a target for the policy instrument. Moreover, in both cases the LR test for the restriction is around the 5% critical value.

The above results suggest an important role for the bond rate in monetary policy. And since the inflation term has the wrong sign in the unrestricted relation under \mathcal{H}_0 , it is natural to impose a zero restriction on β_4 . Imposing this restriction yields the results reported under \mathcal{H}_2 . In both deterministic specifications, the restriction is clearly accepted. Without the inflation term, the long-run relation is close to being formulated in the interest rate spread; which also implies a simple interpretation of the theoretical relation in (4).

Under \mathcal{H}_3 we have reported the results after imposing the additional restriction, $\beta_2 = -1$. The restriction produces LR statistics corresponding to p -values of 0.55 and 0.37, for the two deterministic specifications. In this model the coefficient to the unemployment

	(A) Fifth order polynomial in u_t									
	\mathcal{H}_0		\mathcal{H}_1		\mathcal{H}_2		\mathcal{H}_3		\mathcal{H}_4	
	α	β^*	α	β^*	α	β^*	α	β^*	α	β^*
f_t	-0.049 (-3.65)	1	-0.007 (-0.76)	1	-0.039 (-2.46)	1	-0.039 (-2.45)	1	-0.078 (-4.48)	1
b_t	-0.014 (-0.68)	-1.287 (-4.93)	-0.033 (-2.51)	0	-0.017 (-0.72)	-1.010 (-11.67)	-0.018 (-0.75)	-1	0	-1
$u_t - u_t^*$	-0.007 (-0.66)	2.945 (10.46)	0.000 (0.01)	4.752 (11.83)	-0.032 (-2.49)	3.006 (13.98)	-0.032 (-2.48)	3.008 (14.24)	0	2.454 (13.98)
π_t	-0.054 (-4.86)	1.152 (3.55)	-0.039 (-5.36)	1.239 (4.33)	-0.060 (-4.63)	0	-0.061 (-4.63)	0	-0.057 (-3.64)	0
1		-18.628 (-7.60)		-39.786 (-12.08)		-16.025 (-10.90)		-16.107 (-13.03)		-13.378 (-12.62)
Loglik	1407.384		1405.211		1406.780		1406.780		1404.083	
LR statistic	...		4.346		1.208		1.208		6.602	
p -value	...		[0.04]		[0.27]		[0.55]		[0.16]	
	...		$\chi^2(1)$		$\chi^2(1)$		$\chi^2(2)$		$\chi^2(4)$	

	(B) Fifth order polynomial in u_t and linear trends in $(f_t : b_t : \pi_t)$									
	\mathcal{H}_0		\mathcal{H}_1		\mathcal{H}_2		\mathcal{H}_3		\mathcal{H}_4	
	α	β^*	α	β^*	α	β^*	α	β^*	α	β^*
f_t	-0.054 (-4.36)	1	0.020 (2.79)	1	-0.044 (-2.81)	1	-0.046 (-2.76)	1	-0.083 (-4.77)	1
b_t	-0.010 (-0.53)	-1.380 (-4.60)	0.002 (0.14)	0	-0.005 (-0.23)	-1.119 (-5.15)	-0.013 (-0.53)	-1	0	-1
$u_t - u_t^*$	0.000 (0.01)	2.746 (9.72)	0.036 (-6.34)	2.783 (14.97)	-0.032 (-2.54)	2.888 (12.42)	-0.033 (-2.46)	2.904 (15.70)	0	2.373 (18.07)
π_t	-0.046 (-4.38)	1.576 (4.02)	0.004 (0.63)	-4.818 (-8.04)	-0.055 (-4.26)	0	-0.060 (-4.37)	0	-0.052 (-3.30)	0
1		-19.439 (-5.36)		2.682 (1.01)		-15.125 (-5.19)		-16.083 (-15.57)		-13.347 (-19.09)
Loglik	1408.882		1407.118		1407.911		1407.888		1405.525	
LR statistic	...		3.528		1.942		1.988		6.714	
p -value	...		[0.06]		[0.16]		[0.37]		[0.15]	
	...		$\chi^2(1)$		$\chi^2(1)$		$\chi^2(2)$		$\chi^2(4)$	

Table 3: Testing hypotheses on the long-run structure. t -values based on asymptotic standard errors in parentheses. $\beta^* = (\beta' : \beta_0')'$ denotes the augmented cointegration vector.

gap is around three, and the Federal funds rate and inflation adjust to deviations from the long-run relation. The bond rate seems exogenous; and the adjustment coefficient to the unemployment gap has an unexpected negative sign, indicating that a high interest rate decrease unemployment.

Imposing the additional restrictions that b_t and $u_t - u_t^*$ are weakly exogenous for the long-run coefficients, $\alpha_2 = \alpha_3 = 0$, produces the preferred results reported under \mathcal{H}_4 . The preferred structure is accepted as a reduction of the unrestricted model with insignificant $\chi^2(4)$ LR statistics of 6.60 and 6.71.

5.3 CHARACTERIZING US MONETARY POLICY 1988 : 1 – 2004 : 8

In this section, implications of our empirical findings for US monetary policy 1988 : 1 – 2002 : 12 are presented, and the more recent observations, 2003 : 1 – 2004 : 8, are used for post-sample analysis.

Under the preferred restrictions in \mathcal{H}_4 , the results are almost identical for the two deterministic specifications. For the preferred specification without linear trends, the long-run relation can be written as

$$f_t - b_t = -2.454 \cdot (u_t - u_t^*) + 13.377, \quad (10)$$

which is a Taylor-type target for the interest rate spread with a significant impact from the unemployment gap and a zero impact from inflation beyond that contained in the expected inflation via b_t . Deviations from this relation are corrected primarily by the Federal funds rate, f_t , eliminating 8% of a misalignment each month. This may seem like a slow adjustment, but recall that the target is estimated within a dynamic framework where f_t also adjust to news in terms of the first differences. There is also a negative coefficient in the equation for $\Delta\pi_t$, indicating that a high funds rate suppress inflation.

The Federal funds rate and the estimated long-run target are depicted in Graph (A) of Figure 2, where the relationship is extended out of sample to 2004 : 8. The graph clearly demonstrates that the long-run target has been leading the Federal funds rate when major changes in the latter has occurred corresponding to visual evidence of the endogeneity of the Federal funds rate in the system. It is also clear that the static target is relatively volatile which explains the need for interest rate smoothing. In particular, f_t does not jump instantaneously to the estimated target value, but deviations from the relation set out in (10) is a good indicator of the direction of movements in the funds rate. The deviations show that the interest rate in the initial period and during 1988 was lower than suggested by the relationship. A common interpretation relates this to concerns for the financial stability after the crash in the stock market in late 1987. The same type of concern might explain the relatively low interest rates in 1999 after the financial crisis in Russia and the problems related to LTCM. Such effects are clearly outside the information set of the current simple model.

The estimated natural rate of unemployment, u_t^* , interpretable as the rate of unemployment that leaves monetary policy unaffected, is reported in graph (B). We note that the level of u_t^* is not identified in the model, and the graph is constructed by assuming that the natural rate has the same sample average as u_t . The main picture is that u_t^* has been by and large constant until 1994; whereas a considerable proportion of the unemployment fall during 1994 – 2000 is attributed to a fall in u_t^* . We note, that the estimated polynomial $u_t^* = g(t)$, which approximates the natural rate in the estimation is not useful for predictions, and for the out-of-sample analysis, we take a simple approach and assume that u_t^* is constant.

As mentioned above, actual inflation is not directly present in the empirical rule under

\mathcal{H}_4 . An interpretation of this is, that the period under consideration has been characterized by only moderate inflationary pressure and therefore limited information on the impact of actual inflation in the policy rule. Although actual inflation does not enter expected inflation is present with a coefficient of 1 via the long-term interest rate. It is probably a specific feature for US monetary policy that the long-run policy rule can be reduced to a relationship between the interest rate spread and unemployment. The results reported in Christensen (2002) do not suggest that similar policy rules have been in place in UK or Sweden, nor at the ECB.

The fact that the bond rate and unemployment gap are weakly exogenous implies that α_{\perp} will contain unit vectors corresponding to these variables. The interpretation is that the cumulated shocks to these equations constitute driving stochastic trends in the system. The estimate of α_{\perp} is given by

$$\hat{\alpha}_{\perp} = \begin{pmatrix} 0 & 0 & 1 \\ 1 & 0 & 0 \\ 0 & 1 & 0 \\ 0 & 0 & -1.367 \end{pmatrix},$$

indicating that the last driving trend is the cumulated shocks in the equation for $f_t - 1.367\pi_t$, which is close to the shocks to the real funds rate. The restriction that the last driving trend is the shock to $f_t - \pi_t$ corresponds to the restriction that the adjustment coefficient is identical in the equations for Δf_t and $\Delta \pi_t$. This restriction can be accepted with a marginal LR statistic of 0.43 and a p -value of 0.51 in a $\chi^2(1)$ distribution. The non-stationarity thus seems to be driven by business-cycle shocks to the unemployment gap, $u_t - u_t^*$, shocks to the financial markets as measured by b_t , and real monetary policy shocks to $f_t - \pi_t$.

Based on the moving average representation in (6) we can also test if objectives of monetary policy are actually controllable by the Federal funds rate in the estimated system. The long-run impact matrix is given by

$$\hat{C} = \begin{pmatrix} 0.240 & 2.226 & -3.284 & -0.328 \\ (0.25) & (2.87) & (-2.53) & (-0.74) \\ -0.070 & 1.303 & -0.154 & 0.096 \\ (-0.17) & (3.86) & (-0.27) & (0.49) \\ -0.126 & -0.376 & 1.275 & 0.173 \\ (-0.47) & (-1.70) & (3.45) & (1.36) \\ -0.494 & 0.487 & -0.727 & 0.675 \\ (-2.14) & (2.57) & (-2.30) & (6.20) \end{pmatrix},$$

with asymptotic t -ratios in parentheses. Recall that the necessary condition for controllability of inflation is that $C_{41} \neq 0$. The relevant coefficient is -0.494 , indicating that a one percentage point innovation to the Federal funds rate, interpretable as an unexpected policy change, lowers the long-run core inflation rate with 0.5 percentage points. According to the t -ratio the coefficient is significant, suggesting that the FOMC can actually control inflation in the present context. For the unemployment gap, the situation is the opposite, with a t -ratio of -0.47 .

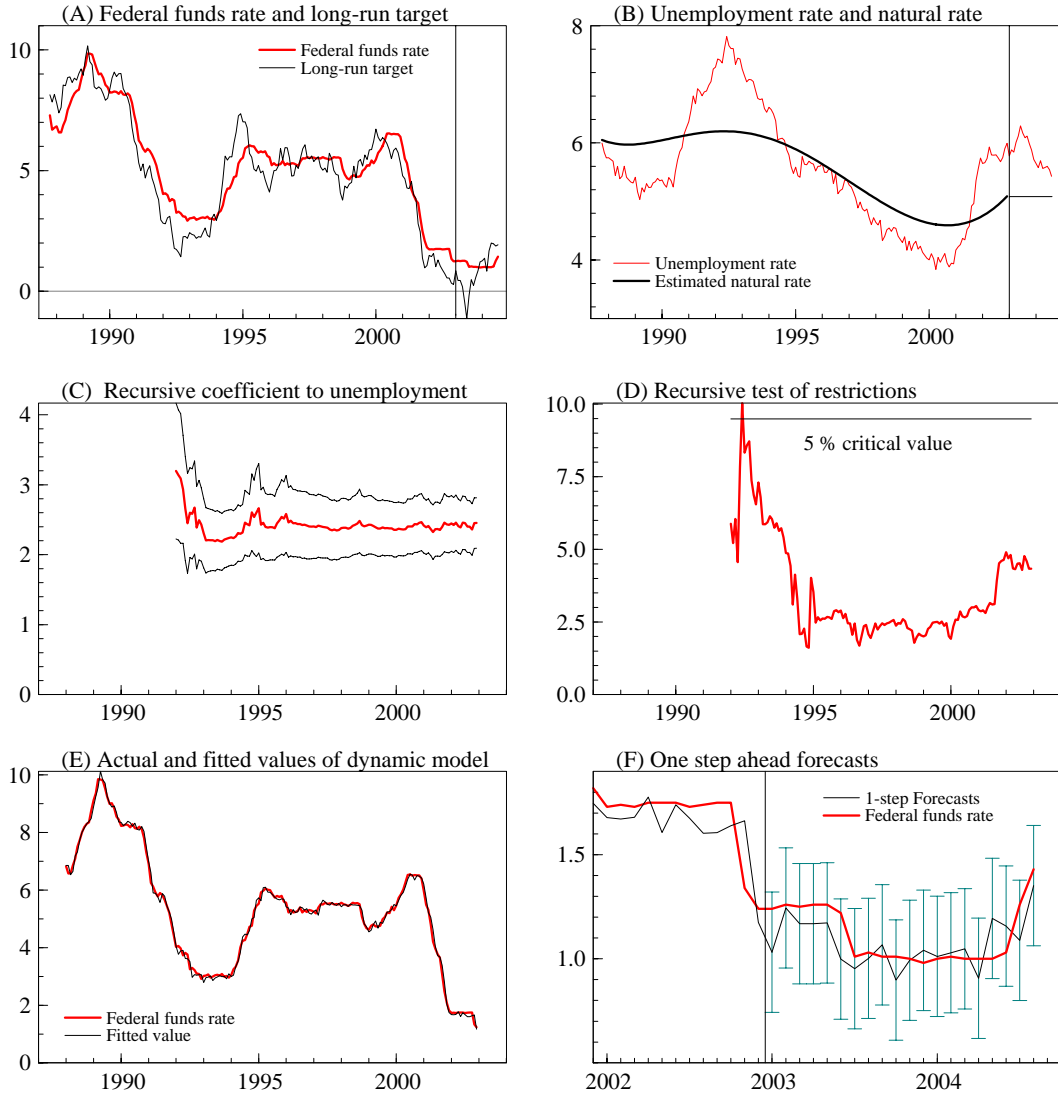


Figure 2: Graph (A) depicts the long-run Taylor-type target and the actual Federal funds rate. Graph (B) depicts the unemployment rate and the estimated natural rate of unemployment, u_t^* . In both cases a vertical line indicates the end of the estimation sample, 2002 : 12. Graph (C) and (D) report results from recursive estimation done for sub-samples $t = 1988 : 1, \dots, T_0$, where the endpoint takes the values $T_0 = 1992 : 1, \dots, 2002 : 12$. In each sub-sample the short-run parameters are fixed at their full-sample estimates, see Hansen and Johansen (1999). By and large similar results are obtained if the short-run parameters are reestimated in each sub-sample, although a larger initial sample is necessary. (C) depicts the recursively estimated coefficient to unemployment in the long-run relation under \mathcal{H}_4 together with the 95% confidence band, while (D) depicts the test statistics for the 4 over-identifying restrictions in \mathcal{H}_4 and the 5% critical value for individual tests, see Kongsted (1998). Graph (E) depicts the actual and fitted values of the Federal funds rate in the dynamic system. Finally, graph (F) depicts the one-step forecasts and the actual Federal funds rate.

5.4 STABILITY OF THE RESULTS

An important issue is the structural stability of the estimates. According to the Lucas-critique view on empirical analyses, only *deep parameters*, such as characterizations of preferences and technical relationships, can be expected to be stable, whereas reaction functions and reduced forms are prone to instabilities following shocks to the system.

To evaluate the stability of the relation we depict in graph (C) the recursively estimated parameters to unemployment in the long-run relation, see Hansen and Johansen (1999). The estimates look remarkably stable, and the narrowing of the 95% confidence bands indicates an increasing information on the parameters. Finally graph (D) depicts the recursively calculated test statistic for the over-identifying restrictions. Apart from a few observations in the beginning of the recursive estimations, where the number of degrees of freedom is small, the identifying structure is clearly acceptable in all sub-samples.

5.5 SHORT-RUN STRUCTURE

To illustrate the dynamic reaction function we apply a general-to-specific modelling strategy and find a parsimonious representation of the system, see Hendry and Mizon (1993). Using a conventional 5% critical level yields the following specification

$$\begin{pmatrix} \Delta f_t \\ \Delta b_t \\ \Delta(u_t - u_t^*) \\ \Delta \pi_t \end{pmatrix} = \begin{pmatrix} -0.080 & 0.248 & 0.154 & -0.200 & 0.173 & 0 \\ (-4.88) & (3.76) & (3.12) & (-2.27) & (2.83) & \\ 0 & 0 & 0.319 & 0 & 0 & 0 \\ & & (4.54) & & & \\ 0 & -0.203 & 0 & -0.188 & 0 & 0.163 \\ & (-4.55) & & (-2.56) & & (2.44) \\ -0.041 & 0 & 0 & 0 & 0 & -0.231 \\ (-3.72) & & & & & (-3.23) \end{pmatrix} \begin{pmatrix} \frac{f_{t-1} - f_{t-1}^*}{\Delta f_{t-1}} \\ \Delta b_{t-1} \\ \frac{\Delta(u_{t-1} - u_{t-1}^*)}{\Delta f_{t-2}} \\ \Delta \pi_{t-2} \end{pmatrix},$$

which produces a LR test statistic of 21.60 compared to the model under \mathcal{H}_4 , corresponding to p -value of 0.54 in a $\chi^2(23)$.

The equation for the Federal funds rate represents the dynamic policy reaction function formulated in terms of observable quantities, and it indicates a simple behavior. Apart from the clearly significant adjustment towards the long-run target, there are only four significant coefficients. Together with the equilibrium correction coefficient, α_1 , the coefficients to the autoregressive terms, Δf_{t-1} and Δf_{t-2} , describe the interest rate smoothing. Besides this feature, there are additional short-run terms to the one period lagged bond rate, Δb_{t-1} , and unemployment rate, Δu_{t-1} , both with short-run coefficients well below the long-run impacts. We can add that the cross-correlations of the system residuals are low, and none of them corresponds to significant contemporaneous effects. The latter finding is most likely a consequence of the time lag in the information set available for monetary policy.

To illustrate the overall fit of the model, graph (E) depicts the actual and fitted values for the funds rate from the dynamic model. The excellent in-sample fit translates to a reasonable out-of-sample behavior. Graph (F) reports the one-step forecasts. Based on

the current information set, the model is not able to predict the last fall in the funds rate in the estimation sample, 2002 : 11, but it is notable that the model is actually able to indicate the direction of the two interest changes in the out-of-sample period, the fall in 2003 : 7 and the rise in 2004 : 7. Both the fitted values and the forecasts are a bit more volatile than actual interest rates, but the main tendency is captured. We can add that the forecast performance for the remaining variables are much weaker, which just reflects that the current information set is best suited for a description of monetary policy.

5.6 INTERPRETATION OF A SECOND LONG-RUN RELATION

In the rank determination in Table 2, the second long-run relation was formally rejected, but close to the 95% quantile of the simulated distribution. In this section we briefly account for a possible interpretation of the second long-run relation, and illustrate that the main findings in the analysis of the case $r = 1$ carry over to the alternative case $r = 2$. In the presentation below, we focus on the case without deterministic linear trends in f_t , b_t , and π_t .

Allowing for a second long-run relation and imposing restrictions on α and β^* yields the structure

$$\begin{pmatrix} \Delta f_t \\ \Delta b_t \\ \Delta(u_t - u_t^*) \\ \Delta \pi_t \end{pmatrix} = \begin{pmatrix} -0.085 & -0.069 \\ (-4.95) & (-4.29) \\ 0 & 0 \\ (...) & (...) \\ 0 & 0 \\ (...) & (...) \\ -0.039 & -0.064 \\ (-2.63) & (-4.55) \end{pmatrix} \begin{pmatrix} f_{t-1} - b_{t-1} + 0.887 \\ (2.71) \\ \pi_{t-1} + 2.475(u_{t-1} - u_{t-1}^*) - 18.539 \\ (15.98) & (-17.88) \end{pmatrix} + \dots,$$

which is not rejected with a test statistic of 11.72 in a $\chi^2(7)$ distribution. Note that for $r = 2$, the cointegration space separates into a stationary interest rate spread, $f_t - b_t$, and a stationary relation between inflation, π_t , and the unemployment gap, $u_t - u_t^*$, interpretable as a Phillips curve relation. The policy target is therefore no longer explicit as a long-term relationship, but the main conclusions from the analysis of $r = 1$ are preserved:

The Federal funds rate, f_t , significantly equilibrium corrects to both long-run relations, and based on the weights of the two relations in the dynamic equation for Δf_t , the implicit policy target has a coefficient to the unemployment gap of around 2.0. Under this structure there is a small negative coefficient to the actual inflation in the policy reaction function, as it was also found in the unrestricted structure for the case $r = 1$. The bond rate and the unemployment gap are still weakly exogenous for the long-run parameters, while inflation, π_t , equilibrium corrects to both relations.

6 CONCLUDING REMARKS

In this paper we reconsider the empirical relationships between the Federal funds rate, unemployment, inflation, and the long-term government-bond rate by means of cointe-

gration techniques. As a starting point, we reinterpret the simple Taylor-type policy rule as a longer term target for the policy instrument, which makes the actual policy reaction function a dynamic adjustment towards the target. Being formulated in terms of observables, the policy target inherits any volatility of the observed data and it is natural that interest rate smoothing should be an integral part of the description of monetary policy.

As the target for the Federal funds rate we find a simple, yet stable, relationship comprising the bond rate and the deviation of the unemployment rate from an estimated natural rate. The estimated relation is interpretable as a monetary policy target because the Federal funds rate seems to be the main variable correcting deviations from the equilibrium relationship. The presence of the bond rate is interpreted as indicating an important role of financial market information, most notable the information on expected inflation contained in bond yields. The result also fits into recent work on central banking as managing expectations as a supplement to setting short-term interest rates. The objectives of monetary policy and the FOMC's assessment of the stance of the economy are clarified in statements and testimonies, and the financial market may react before policy changes are actually implemented in the Federal funds rate. Within the estimated system inflation *is* controllable with the Federal funds rate, but since the registered unemployment rate is uncontrollable, the main transmission is *not* via the labour market. The implications of this finding are left for future research.

The real-time monetary policy process is—without doubt—far more complicated than indicated by the estimation results obtained for the current information set. We believe, however, that the present study gives a theoretically understandable and empirically relevant characterization of the monetary policy process in terms of observable variables. Moreover, the interpretation of interest rate smoothing as equilibrium correction ensures that the variables in the estimated target are actually driving variables for the Federal funds rate in the empirical model, and not just covariates in a claimed structural policy relation. A traditional Taylor-rule, where the Federal funds rate reacts to inflation and the unemployment gap, is clearly rejected.

Naturally, it would be interesting to reconsider the results for the policy process based on richer information set; and it would also be of interest to consider other time periods and countries. These extensions are, however, far beyond the scope of the present paper.

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